The Effect of Foreign Direct Investment on Urban Wage: An Empirical Examination

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Abstract:

How does the openness to foreign direct investment (FDI) affect the wage level of cities? In this paper we use a panel data set of Chinese cities to examine the impact of inward FDI on urban real wage. The results suggest that the existence of FDI has a significant and positive effect on urban real wage and this impact remains significant after controlling other city characteristics.

JEL Classification: F43; J31; O18

1. Introduction

Foreign direct investment (FDI) plays an increasing important role in global economy. Global flows of FDI have grown rapidly during the last decade and FDI becomes the largest source of foreign private capital flow into the developing countries. The policy to attract FDI is becoming one of important economic development strategies for developing countries, based on the assumption that FDI has beneficial host country effect. Evaluation of this policy requires a comprehensive study of the host country effect of inward FDI.

One of central questions of this study is how FDI affects the wage level in the host country. To answer this question, previous studies have identified several mechanisms through which FDI may have different impact on wages in host countries. Brown, Deardorff and Stern (2003) provided a theoretical overview on these mechanisms, and they concluded: "All of the cases we have considered in this theoretical overview – capital flow, technology flow, and fragmentation – have failed to yielded unambiguous conclusions about the effects of FDI and multinational firms on equilibrium wages in host countries. … It is therefore an empirical question whether the actual operations of multinationals have raised or lowered wages in developing countries." (Brown, Deardorff and Stern, 2003, p. 40)

Most empirical studies on this topic used plant-level data to investigate two questions: first, do foreign-owned firms pay higher wage than local firms? Second, is there wage spillover effect from foreign-owned firms to local-owned firms? Brown, Deardorff and Stern (2003) and Lipsey (2004) provided nice surveys of these empirical evidences. However, only few studies focus on the aggregate impact of FDI on regional average wage level. As Lipsey (2004) pointed out, a policy maker may be more interested in the aggregate impact of FDI on regional average wage level, no matter which sources this impact comes from. Economic theories can not provide unambiguous answer to this question and it calls for the further empirical examination. This paper is, to the best of our knowledge, the first in directly testing the aggregate effect of inward FDI on the wage level of cities.

This paper uses the information of Chinese cities to investigate the impact of city openness to FDI on city wage level. We focus on the case of China for several reasons: first, since 1990, China has experience high growth rate in inward FDI flow. China becomes the largest recipient of FDI among developing countries and the second largest FDI recipient in the world[1]. Chinese experience may provide helpful insights to the study on host country effect of FDI. Second, compared to the cross-country study, the regional study within the same country may have advantage in controlling regional heterogeneity such as political, institutional and cultural factors. Third, the geographic distribution of FDI across regions within China is extremely uneven, and thus provides a good opportunity to test the impact of FDI on regional wage level.

In this study we estimate the wage equation using the system Generalized Method of Moments (GMM) estimator. The results suggest that the existence of FDI has a significant and positive impact on urban real wage, and this impact remain significant after controlling other city characteristics. Cross-city variation of the openness to FDI is one of important contributing factors of intercity wage inequality.

The remainder of this paper is organized as follows: Section 2 briefly reviews the literature. Section 3 describes the data set. Section 4 investigates the impact of openness to FDI on urban real wage. Section 5 concludes this paper.

2. Literature

How do foreign direct investments affect the wage level in the host country? In their theoretical overview, Brown, Deardorff and Stern (2003) considered four roles of FDI, each of which may have different implication for wages in host country. First, FDI may carry additional capital into the host country. In the normal case, the increased capital stock will raise the marginal product of labour and thus its wage in competitive factor markets. But there are several cases where wages do not change or even fall. Second, FDI may carry superior technology into host country. The effect of this improvement of technology on wages is not unambiguous and depends on the circumstances. Third, multinational productions may involve the production processes that are fragmented across countries. The effect of this fragmentation on local factor prices depends on factor intensity of fragments relative to factor endowments of the country. Fourth, multinational firms may have some market power to set prices in local labour market, which is higher or lower than local firms. Multinational firms may tend to pay above-equilibrium wages. Lipsey (2004) summarized several possible explanations: the foreign firms may be forced to offer higher wage by host country regulation or pressure. It’s also possible that the foreign firms have to compensate the local labour to overcome their preference to home firms. Because of the information disadvantage in local labour market, the foreign firms may offer higher wage to attract better workers. The foreign firms may also use this wage premium to reduce worker turnover and slow down the technology leak. Brown, Deardorff and Stern (2003) concluded that theoretical analysis does not provide unambiguous prediction on the effect of FDI on wages in host country.

Previous empirical studies mainly focus on two effects of FDI on wages: the wage differential effect and the wage spillover effect. First, the foreign owned firms may offer higher wage than the domestic firms. A large body of empirical literature tried to test this wage differential effect and to identify the sources. Brown, Deardorff and Stern (2003) and Lipsey (2004) surveyed the empirical evidences and concluded that there are overwhelming evidences supporting that foreign firms pay higher wage than domestic firms, even after controlling for scale, worker quality, industry and other industry and regional characteristics.

Second, the higher wages paid by the foreign firms may also results in the higher wages in the domestic firms. This wage spillovers effect may come from increasing competition in the local labour market between foreign and domestic firms, or from knowledge spillovers for foreign firms to domestic firms. Aitken, Harrison and Lipsey (1996) found the lack of wage spillovers but significant wage differentials between foreign and domestic firms in Mexico and Venezuela. In the United States, there is small wage differential but some evidences supporting the wage spillovers effect. Feliciano and Lipsey (1999) found little wage spillovers effect in American manufacturing industries but large and significant effect in non-manufacturing industries. Girma, Greenaway and Wakelin (2001) studied the FDI in UK and found no overall wage spillover effect on wage level but a negative effect on wage growth.

There are only few empirical evidences on the aggregate effect of FDI on regional average wage levels. Aitken, Harrison and Lipsey (1996) found that the foreign ownership shares tended to raised average industry wages in Mexico and Venezuela. Figlio and Blonigeb (2000) studied county-industry data from South Carolina and found that the foreign investment raised local industry wage much more than does domestic investment. Our study extends this line of literature in the following ways: first, we provide the Chinese experience, which is important in the study of host effect of FDI; second, we direct test the impact of FDI on the urban real wage instead of industry average wage; third, since the wage evolution shows high autoregressive property, we use system GMM to estimate the dynamic panel model.

3. Data and Summary Statistics

Our sample includes 231 prefecture-level cities in China during the period 1990 to 1998. The information of city characteristics is mainly compiled from *Cities China, 1949-1998* (State Statistical Bureau, 1999). For these cities, information of both "Shiqu" (urban area) and "Diqu" (urban area and rural counties) are reported, and we only use the information of urban area. Since the information of consumer price index (CPI) in each city is not available, the relevant provincial annual CPI is used to deflate nominal wage into 1990 constant prices. This deflator has limitations since it omits the price variation across different cities within the same province. The information of provincial CPI is compiled from *Comprehensive Statistical Data and Materials on 50 Years of New China* (State Statistical Bureau, 1999). The official exchange rate is used to change the volume of foreign direct investment from US dollar into Chinese currency, and the information is compiled from *Chinese Statistical Yearbooks* (State Statistical Bureau, 2000). The information of education levels of Chinese cities in 1990 is compiled from *Population Statistical Yearbook of China* (State Statistical Bureau, 2000).

In this study we consider two types of city characteristics: the characteristic in our central interest is the city openness to FDI. We use FDI intensity instead of the absolute volume of FDI to measure the openness to FDI. Two alternative measures of FDI intensity are used in this study: one is the ratio of FDI to GDP, which captures the importance of FDI in city economy, and another is the ratio of FDI to the total fixed asset investment FDI intensity, which captures the importance of foreign investment relative to domestic investment. We also consider other city characteristics that may closely link with the wage level (i.e., capital labour ratio, industry structure, and human capital stock).

[Table 1 about here]

The summary statistics of various city characteristics are reported in Table 1. The sample average of the logarithm of real wage is 7.89 while the standard deviation is 0.28. The mean of the share of FDI in city GDP is 0.05, and the standard deviation is almost 2 times the mean. The sample mean of the ratio of FDI to fixed asset investment is 0.20 while the standard deviation is 0.34. This implies high cross-city variation of FDI intensity. The mean of logarithm of total fixed asset investment per employee is about 7.98. The average ratio of employment in the manufacturing sector relative to the service sector is about 1.48. The sample average is 53% for the fraction of people with at least 9 year education, and 22% for the fraction of people with at least 12 year education.

[Figure 1 about here]

The substantial standard deviation of FDI intensity implies large cross-city variation of openness to FDI in China. We show the time trend of Gini coefficients from 1990 to 1998 in Figure 1. For the logarithm of foreign direct investment flow, the Gini coefficient decreased from 0.17 to about 0.1 during the period of 1990-1993, and stayed around 0.1 from 1993 to 1998. The cross city variation of FDI intensity is much higher than the one of FDI volume. For both measures of FDI intensity, the Gini coefficient is extremely high, ranged from 0.59 to 0.78, which implies substantial variation of FDI intensity across cities.

4 The Impact of FDI on Urban Wage Level

4.1 Specification and Methodology

To investigate the impact of the FDI on urban average wage level, we specify the following dynamic panel model:

$$Log(W_{i,t}) = \alpha + (1-\lambda)Log(W_{i,t-1}) + \sum_{i=1}^{T} \beta_j FDI\,intensity_{i,t-j} + X'_{i,t}\delta + \eta_t + \mu_i + \varepsilon_{it} \qquad (1)$$

Where $W_{i,t}$ is the real wage level (in 1990 constant price) of city i in time t. There are two measures of $FDI\,intensity_{i,t}$: one is $\frac{FDI_{i,t}}{K_{i,t}}$, defined as the ratio of FDI to total fixed asset investment of city i in the time t; another one is $\frac{FDI_{i,t}}{GDP_{i,t}}$, defined as the ratio of FDI to GDP of city i in time t. The vector $X_{i,t}$ include other city characteristics which may affect urban wage level; α is a constant; η_t is a set of time dummies; μ_i is the individual effect of city i; and $\varepsilon_{i,t}$ is error term. λ, β and δ are estimated coefficients. In our study β is the coefficient in the central interest, which indicates the impact of lagged FDI intensity on average real wage.

If the observed correlation between FDI and wage is driven by some variables omitted from the regression, then this causal linkage might be misinterpreted. To address this issue, we try to control the city characteristics which possibly affect the wage level. The first control variable is $Log(K_{it}/L_{it})$, defined as the logarithm of total fixed asset investment per employee. Higher capital labour ratio may increase the marginal productivity of labour and thus the wage payment. The second variable is the measure of industrial structure of city. With significant wage differential across industries, the different industrial mix may be an important source of inter-city wage inequality. Unfortunately the detail information of industrial structure of Chinese cities is not available in our data set. We use $(L_{it}^{Manufacture}/L_{i,t}^{Service})$, the ratio of employment in manufacturing sector relative to the employment in service sector, as the indicator of sectoral composition. In the urbanization and industrialization process, employment in the agricultural sector shifts toward the manufacturing sector, and then toward the service sector. The service industries may be more skill intensive than the manufacturing industries, and thus service-based cities may have higher wage level than manufacturing-based cities. The third variable is the human capital accumulation. High human capital accumulation contributes to the marginal labour productivity, and thus yields higher wage payment. We use two standard measures of human capital stock: $Share^{9years}$, defined as the fraction of city population with at least 9 years education, and $Share^{12years}$, defined as the fraction of city population with at least 12 years education.[2] Due to data limitation, we are not able to control cross-city variation of unemployment and infrastructure investment. The summary statistics of these controlling variables are reported in Table 1 and discussed in Section 2.

For this dynamic panel model, the OLS estimator and random effect estimator are biased since lagged dependent variable is included as a regressor and is correlated with the individual effect μ_i. Although the fixed-effect estimator eliminates the individual effect by transforming data into deviations from the within-group mean, it is still biased because the within-group mean of lagged wage is correlated with the mean of the error term. The suitable method to estimate this model is the Generalized Method of Moments (GMM). Arellano and Bond (1991) developed the first-differenced GMM estimator. This method first differences the data to eliminate the individual effect, μ_i, and then utilizes all the lagged values of the regressors as instruments. However, when the regressors are highly persistent, or close to random walk, the lagged levels of regressor are only weak instruments for the first differences of the series, and thus the first-difference GMM estimator has poor finite sample property. Arellano and Bover (1995) and Blundell and Bond (1998) showed that the lagged first differences of the variables may be suitable instruments for the equation in levels, and the so-called system GMM estimator was developed by combining both sets of moment conditions in a system containing both first-differenced and level equations. Our study adopts the system GMM estimator, which has been shown to have large efficiency gain over the first-difference GMM estimator. We also use a finite-sample correction to the two-step covariance matrix (Windmeijer, 2000) to obtain consistent estimates of standard error.

4.2 Results

The wage equation (1) is estimated using the two–step system GMM estimator and the results are reported in Table 2. Column 1, 2, 7 and 8 of Table 2 report the system GMM estimates of the lagged real wage level and lagged FDI intensity, without controlling other urban characteristics. Time dummies are included to control for the possible variation of macroeconomic environment over time.

[Table 2 about here]

As Table 2 show, the estimated coefficient of the lagged real wage level is significant and positive, with $\hat{\lambda}$ close to 0.1. This shows high autoregression dynamic of city average wage level. The estimated coefficients of one-year lagged FDI intensity are significant and positive: a 1% increase in the share of FDI in total fixed asset investment may lead to 0.055% increase in the logarithm of real wage, and a 1% increase in the share of FDI in city GDP may lead to 0.185% increase in the logarithm of real wage. The presence of FDI tends to raise the average urban wage level. a Hansen's test of overidentificantion restriction (Hansen, 1982) doesn't reject the null hypothesis, and verifies the validity of the moment conditions.[3] Arellano-Bond tests for autocorrelation (Arellano-Bond, 1991) verify the assumption that there is first order but no second order autocorrelation.

After both one-year and two-year lagged FDI intensity are included into the regression, the estimated FDI-wage linkage becomes insignificant. For the ratio of FDI to total fixed asset investment, the validity of model is rejected by the Hansen's J statistics at the significance level of 1%. This is possibly due to multicollinearity problem since the one-year lagged FDI intensity and two-year lagged FDI intensity are highly correlated. The Pearson correlation coefficient between one-year and two year lagged ratio of FDI to the GDP is 0.78; and the correlation between one-year and two-year lagged ratio of FDI to total fixed asset investment is 0.68. To avoid this multicollinearity problem, we only include the one year lagged FDI intensity in the regression.

There is possible endogeneity bias in this study. City wage level may affect the location choice of FDI. To minimize the cost, foreign investors will choose to invest in low wage region if other regional characteristics are identical. For example, Sun, Tong and Yu (2002) found that lagged provincial wage level had a negative impact on provincial FDI flow in the 1990s. This negative linkage between FDI and wage will yield endogeneity bias and underestimate the positive impact of FDI on urban wage level. To reduce this bias, we use the lagged value of FDI intensity since it is predetermined to the current wage level. The system GMM estimator includes both level and difference of lagged regressors in the instrument matrix and thus control for the possible endogeneity of the explanatory variables.

[Table 3 about here]

For robustness test, we include other city characteristics in the regression to address the omitted variable bias. A potential problem is the multicollinearity among the regressors. To ascertain the degree of multicollinearity, we report the correlation matrix between the regressors in Table 3. As this table shows, the high correlation only exists between two measures of FDI intensity (the correlation coefficient is 0.777), and between the two measures of human capital accumulation (the correlation coefficient is 0.868). Since each of these measures enters the regression separately, multicollinearity is not a serious problem in our study.

Column 3 – 6 and Column 9 – 12 in Table 2 present the estimation results. After controlling other city characteristics, the effect of FDI intensity is still statistically significant and positive. The estimated coefficients of the ratio of FDI to total fixed asset investment are ranged from 0.054 to 0.059. The estimated coefficients of the ratio of FDI to GDP are ranged from 0.173 to 0.207. There is still strong positive effect of lagged wage level, and the estimated coefficients become smaller than the one in the benchmark result (Column 1 and 7 in Table 2). As expected, the estimated coefficient of $Log(K_{it} / L_{it})$ is positive and significant, which implies that the higher physical capital investment contributes to higher aggregate wage level. The estimated coefficient of manufacturing to service employment ratio is negative but statistically insignificant. The measures of human capital accumulation have positive and weakly significant impact on the city wage level. For each regression, Hansen's J statistics verify the validity of the model and the Arellano-Bond tests confirm the assumption that there is first order but not second order autocorrelation.

5 Concluding Remarks

How does inward FDI affect the wage level in the host country? This question is important for the policy makers, especially in developing countries, to evaluate the policies attracting FDI. Theoretical analysis identified several ways how FDI affect the wages in host country, but failed to yield unambiguous conclusion. This paper uses city level data to directly test the aggregate impact of FDI on urban wage level in China. The results show that the openness to FDI tends to increase the average wage level of Chinese cities. This positive impact remains significant after controlling for the city capital-labour ratio, industrial structure and human capital stock. Our results also suggest that both physical capital and human capital accumulation tend to raise the average wage level of cities.

References

Aitken, B., Harrison, A. and Lipsey, R. (1996) Wages and Foreign Ownership: A Comparative Study of Mexico, Venezuela, and the Unites States, **Journal of International Economics***,* 40, pp. 345-371.

Arellano, M. and Bond, S. (1991) Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations, **Review of Economic Studies***,* 58, pp. 277-297.

Arellano, M. and Bover, O. (1995) Another Look at the Instrumental-Variable estimation of Error-Components Models, **Journal of Econometrics***,* 68, pp. 29-52.

Blundell, R. and Bond, S. (1998) Initial Conditions and Moment Restrictions in Dynamic Panel Data Models, **Journal of Econometrics***,* 87, pp. 115-143.

Brown, D., Deardorff, A. and Stern, R., 2003, the Effects of Multinational Production on Wages and Working Conditions in Developing Countries, NBER Working Paper No. 9669.

Feliciano, Z. and Lipsey, R. (1999) Foreign Ownership and Wages in the United States, 1987-1992, NBER Working Paper No. 6923

Figlio, D. and Blonigen, B. (2000) The Effects of Foreign Direct Investment on Local Communities, **Journal of Urban Economics***,* 48, pp. 338-363.

Girma, S., Greenaway, D. and Wakelin, K. (2001) Who Benefits from Foreign Direct Investment in the UK? **Scottish Journal of Political Economy**, 48, pp. 119-133.

Hansen, L. (1982) Large Sample Properties of Generalised Method of Moments Estimators, **Econometrica***,* 50, pp. 1029-1054.

Lipsey, R. (2004) Home and host country effects of FDI, in R. Baldwin and A. Winters (Ed.) **Challenges to Globalization: Analyzing the Economics**, pp. 379-382, Chicago: The University of Chicago Press.

Sargan, J. (1958) The Estimation of Economic Relationships Using Instrumental Variables, **Econometrica**, 26, pp. 393-415.

State Statistical Bureau, (1999) **Cities China, 1949-1998**, Beijing: China Statistics Press.

------- (1999) **Comprehensive Statistical Data and Materials on 50 Years of New China**, Beijing: China Statistic Press.

------- (2000) **Population Statistical Yearbook of China,** Beijing: China Statistics Press.

------- (2000) **Chinese Statistical Yearbook**, Beijing: China Statistics Press.

Sun, Q., Tong, W. and Yu, Q. (2002) Determinants of Foreign Direct Investment Across China, **Journal of International Money and Finance**, 21, pp. 79-113.

Windmeijer, F. (2000) A Finite Sample Correction for the Variance of Linear Two-step GMM Estimators, Working Paper 00/19, Institute for Fiscal Studies, London.

Table 1
Summary of Statistics of Urban Characteristics

	Mean	S.D.	25%	Median	75%	Observatio n
Log (real wage)	7.89	0.28	7.70	7.85	8.04	2017
The ratio of FDI to GDP	0.05	0.09	0.01	0.02	0.06	1747
The ratio of FDI to total fixed asset investment	0.20	0.34	0.03	0.08	0.22	1744
Log(fixed assets investment per employee)	7.98	1.08	7.32	7.97	8.69	2023
The employment ratio of manufacture to service	1.48	0.77	0.98	1.33	1.80	2029
The fraction of people with at least 9 year education	0.53	0.12	0.43	0.57	0.64	2025
The fraction of people with at least 12 year education	0.22	0.09	0.14	0.23	0.28	2025

Table 2
Determinants of Urban Average Wage Level

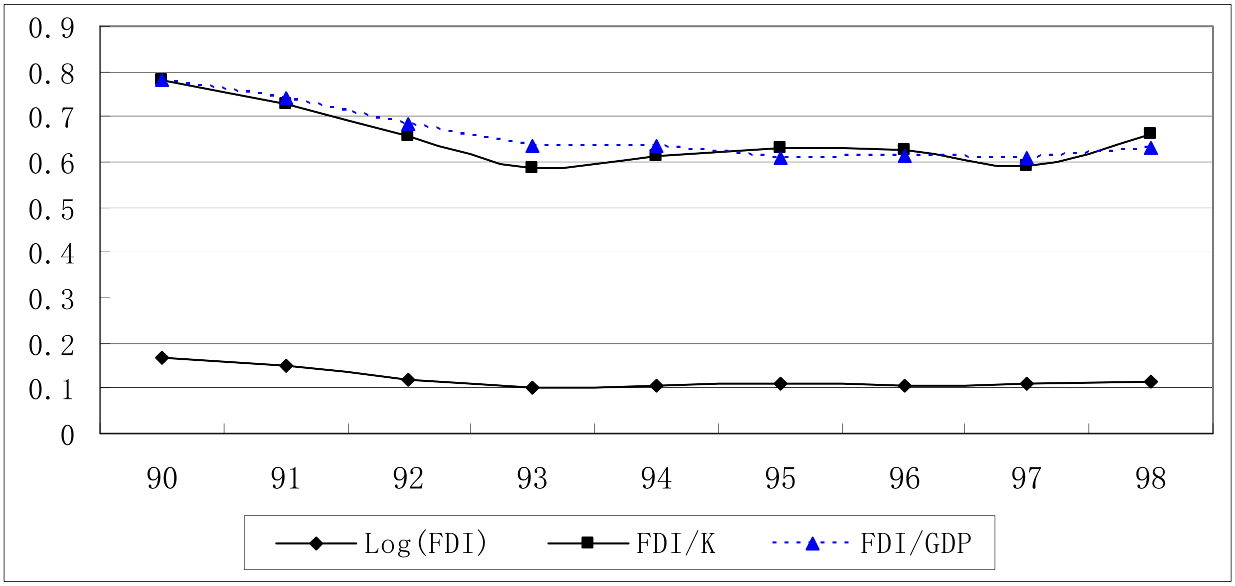
Variables	The Ratio of FDI to Total Fixed Asset Investment						The Ratio of FDI to GDP					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Constant	0.698* (1.93)	1.750** (3.56)	1.124** (2.29)	1.533** (3.12)	1.546** (3.19)	1.710** (3.19)	0.931** (2.13)	0.770** (2.26)	1.672** (3.32)	1.985** (3.95)	1.760** (3.33)	1.832** (3.29)
$Log(W_{it-1})$	0.918** (19.49)	0.783** (12.24)	0.854** (12.21)	0.781** (11.21)	0.757** (10.42)	0.744** (9.39)	0.888** (15.61)	0.910** (20.64)	0.766** (10.47)	0.724** (9.60)	0.731** (9.53)	0.732** (9.07)
$FDI\ intensity_{i,t-1}$	0.055** (2.33)	0.057 (1.11)	0.059** (2.61)	0.054** (2.92)	0.056** (2.47)	0.055** (2.43)	0.185** (2.30)	0.185* (1.82)	0.207** (2.28)	0.179* (1.86)	0.183* (1.71)	0.173* (1.68)
$FDI\ intensity_{i,t-2}$		0.087 (0.91)						-0.037 (-0.44)				
$Log(K_{it} / L_{it})$			0.017 (1.27)	0.038** (2.78)	0.035** (2.34)	0.031** (2.43)			0.033* (1.80)	0.039* (1.95)	0.033** (2.32)	0.030** (2.22)
$\frac{L_{it}^{Manufacture}}{L_{i,t}^{Service}}$				-0.016 (-0.79)	-0.028 (-1.38)	-0.027 (-1.44)				-0.024 (-0.96)	-0.030 (-1.36)	-0.024 (-1.19)
$Share^{9\,years}$					0.251** (1.96)						0.262* (1.89)	
$Share^{12\,years}$						0.416* (1.88)						0.326* (1.67)
P value of Hansen's test of overidentification restriction	0.322	0.000**	0.161	0.134	0.100	0.140	0.479	0.728	0.189	0.217	0.147	0.178
P value of Arellano-Bond test for AR(1)	0.031**	0.040**	0.031**	0.026**	0.026**	0.027**	0.030**	0.042**	0.029**	0.027**	0.027**	0.028**
P value of Arellano-Bond test for AR(2)	0.157	0.130	0.217	0.246	0.216	0.205	0.166	0.084*	0.226	0.192	0.180	0.183
Number of observations (Number of firms)	1510 (224)	1246 (217)	1504 (224)	1504 (224)	1474 (218)	1474 (218)	1513 (224)	1250 (218)	1507 (224)	1507 (224)	1477 (218)	1477 (218)

Note: t-statistics are provided in parenthesis. * and ** represent significance at the 10% and 5% levels respectively.

Table 3
Correlation Matrix

	$\frac{FDI_{i,t-1}}{GDP_{i,t-1}}$	$\frac{FDI_{i,t-1}}{K_{i,t-1}}$	$Log(K_{it}/L_{it})$	$\frac{L_{it}^{Manufacture}}{L_{i,t}^{Service}}$	$Share^{9\,years}$	$Share^{12\,years}$
$\frac{FDI_{i,t-1}}{GDP_{i,t-1}}$	1.000	–	–	–	–	–
$\frac{FDI_{i,t-1}}{K_{i,t-1}}$	0.777	1.000	–	–	–	–
$Log(K_{it}/L_{it})$	0.252	0.156	1.000	–	–	–
$\frac{L_{it}^{Manufacture}}{L_{i,t}^{Service}}$	-0.251	-0.235	0.001	–	–	–
$Share^{9\,years}$	-0.061	-0.078	0.281	0.191	1.000	–
$Share^{12\,years}$	-0.006	-0.037	0.302	0.061	0.868	1.000

Figure 1
Cross City Variation of Openness to FDI, 1990-1998



Note:

[1] Since 2003, China has become the largest FDI recipient in the world.
[2] The information of education structure of city population is only available in 1990, the beginning year of our sample period. With this data limitation, we are not able to control the dynamic change of human capital stock during the sample period
[3] The Sargan statistic (Sargan, 1958) is widely used to test the overidentifying restrictions. This statistic is not robust to heteroskedasticity or autocorrelation. Instead of the Sargan statistic, we report the Hansen J statistic (Hansen, 1982), which is the minimized value of the two-step GMM criterion function, and is robust.